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REVISITING THE IMPLICATIONS OF HETEROGENEITY IN FINANCIAL MARKET PARTICIPATION FOR THE C-CAPM

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Abstract

Recent studies have explored the possibility that accounting for limited participation in financial markets, and in the stock market in particular, might explain the empirical inconsistency of the Consumption-based Capital Asset Pricing Model (C-CAPM). The reasoning is that if non-shareholders' consumption growth co-varies less than shareholders' with share returns, it is misleading to include their expenditure in the consumption measure used to test the model. This paper reviews the implications of household portfolio heterogeneity for various well-known characterizations of the empirical failure of the model, such as the inconsistency of consumption-based asset pricing factors with Hansen and Jagannathan bounds, the equity premium puzzle and the rejection of the overidentifying restrictions to the model. Specifically, it provides a unified framework of analysis, based on the US Consumer Expenditure Survey, to assess the extent to which the empirical inconsistency of the C-CAPM can be attributed to the use of aggregate data that do not allow us to account for limited participation in asset markets. The evidence supports the view that accounting for portfolio heterogeneity improves the empirical performance of the model and helps rationalize some of the puzzling findings, but by itself is not enough to reconcile the theory with the empirical evidence.

JEL classification: E21, G12.

Keywords: Consumption-based Capital Asset Pricing Model, limited participation to financial markets.

Contents

1. Introduction[1]

The Consumption-based Capital Asset Pricing Model (C-CAPM) by Merton (1973), Breeden (1979), Lucas (1978) and Brock (1982) is a popular framework for understanding the valuation of assets and the serial correlation properties of asset returns and consumption. Its development represents an extension of the Capital Asset Pricing Model (CAPM) of Sharpe (1964) and Lintner (1965) in an attempt to improve the empirical performance of equilibrium asset pricing theories. The C-CAPM is based on the standard representative agent model for consumption, and prices assets by discounting returns using the intertemporal marginal rate of substitution. Specifically, the price of a claim is determined by the product of its expected payoff and the marginal rate of substitution of the representative investor. Empirically, the C-CAPM has not performed well. A well-known characterization of its empirical inconsistency is the Mehra and Prescott equity premium puzzle (1985), which refers to the impossibility of finding a plausible pair of values for the subjective discount rate and the coefficient of relative risk aversion consistent with the observed risk premium. In the literature, there is no consensus on a reasonable value for risk aversion, but most economists believe that a relative risk aversion above 10, or even 5, implies totally unrealistic conduct. Another characterization of the "puzzle" is that of Hansen and Jagannathan (1991), who determine a set of bounds on the first two moments of a generic stochastic asset-pricing factor and find that the moments of the intertemporal marginal rate of substitution are inconsistent with those bounds. Further, Hansen and Singleton (1982, 1984) reject the set of overidentifying restrictions to the model, providing additional evidence against the C-CAPM. The empirical inconsistency of the C-CAPM is powerful evidence against the hypotheses of rational expectations and of market efficiency, which lie behind most modern economic and financial models.

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There is now a vast literature aimed at rationalizing this evidence. Epstein and Zin (1989, 1991), Constantinides (1990), Heaton (1995) and Campbell and Cochrane (1995), among others, focus on individual preferences, relax the assumption of time-separable power utility, and consider the implications of generalized expected utility, habit formation and relative consumption effects. Another possibility that has been examined, for example by Heaton and Lucas (1996), is to allow for incomplete markets and trading costs. The strengths and weaknesses of this sort of approaches are reviewed extensively by Kocherlakota (1996).

More recently, starting with Mankiw and Zeldes (1991), there has been growing interest in the implications of limited participation in financial markets,[2] which would make per capita consumption growth a poor proxy for individual consumption growth when pricing assets on the basis of their covariance with the individual intertemporal marginal rate of substitution. This paper focuses on the implications of limited financial market participation. The objective is to review the empirical performance of the C-CAPM, providing a unified framework for the empirical assessment of the role of heterogeneity in participation for the three best known characterizations of the empirical inconsistency of the model, namely Hansen and Jagannathan (1991), Mehra and Prescott (1985) and Hansen and Singleton (1982, 1984). The idea that accounting for portfolio heterogeneity might explain the disparity between predictions and evidence depends on the presumption that there is no theoretical reason for non-shareholders to adjust their consumption growth in response to predictable changes in share returns (unless their income is correlated with share returns). If non-shareholders' consumption co-varies with returns much less than shareholders', then to include the former in the consumption measure for the C-CAPM will lead to inconsistent estimates and to the rejection of the model's predictions. Vissing-Jørgensen (1999) quantifies the bias and shows that, in a representative agent framework, under standard assumptions regarding preferences, returns and markets, if the expenditure of non-shareholders is uncorrelated with asset returns, the risk aversion estimate (the predicted equity premium) is biased upward (downward) by a factor that depends on the fraction of

[2] Limited participation in financial markets implies that markets are incomplete at the maximum extent with some agents not trading at all.

shareholders' expenditure in the aggregate.[3] Hence if, for whatever reason, only a subset of agents do hold shares, it is their per capita consumption growth that should be used for pricing shares and not overall per capita consumption growth as in Mehra and Prescott (1985), among others.

Mankiw and Zeldes (1991) were the first to assess the relevance of this issue empirically. They find that distinguishing between shareholders and non-shareholders helps explain the size of the equity premium. However, they do not fully resolve the equity premium puzzle. Although their estimate of shareholders' risk aversion is substantially lower than that of non-shareholders', it is still too high with respect to the values that are deemed plausible by the literature. Some possible rationalizations of this result can be found in the features of the data set that they use, the US Panel Study of Income Dynamics (PSID), which only measures food expenditure. If food consumption is non-separable from the other components of household expenditure, Mankiw and Zeldes' evidence could be misleading.[4] Further, in the sample they use, asset holdings are reported only in the last wave of the survey, which makes the distinction between shareholders and non-shareholders quite imprecise in the previous waves and, above all, overstates the incidence of share ownership throughout the 1970-1984 period that they analyze. Nevertheless, their findings are substantially confirmed by Brav and Geczy (1995) who perform the same type of analysis using the US Consumer Expenditure Survey, which has more detailed expenditure data and annual information on asset holdings. Both studies rely for the estimation on only one moment restriction, the unconditional Euler equation. Hence, their estimates can be expected to be rather inefficient.[5] Using a conditional version of the model, i.e. adding information about predictable movements in expected consumption growth rates and asset returns,

[3] If, instead of assuming the existence of a representative agent, we aggregate across all agents (including non-shareholders) taking into account cross-sectional heterogeneity in consumption growth rates, the bias depends on the fraction of shareholders in the population.

[4] See Attanasio and Weber (1995) for the issue of preference non-separability over individual consumption items.

[5] Neither Mankiw and Zeldes (1991) nor Brav, Constantinides and Geczy (2002) provide standard errors for their estimate. Vissing-Jørgensen (1999) bootstraps the confidence intervals of these risk aversion estimates based on one moment restriction. The intervals that she determines are extremely wide, more so for non-shareholders than for shareholders.

Attanasio, Banks and Tanner (2002) and Vissing-Jørgensen (2002) obtain efficient and sensible estimates of preference parameters. Attanasio, Banks and Tanner even fail to reject the overidentifying restrictions, providing additional clear-cut support for limited asset market participation theory.

In this paper, I analyze the implications of heterogeneity in financial market participation *vis à vis* the alternative hypothesis of a representative agent who consumes aggregate per-capita expenditure. The first set of tests are based on Hansen and Jagannathan bounds. This test is extremely general since it does not propose and evaluate alternative parametric specifications but simply characterizes the set of stochastic discount factors consistent with observed returns and verifies the consistency of the "heterogeneous" intertemporal marginal rates of substitution with this set. The second type of test is based on the calibration of the C-CAPM and checks the impact of household heterogeneity on the unconditional restrictions implied by the model, in the spirit of Mankiw and Zeldes (1991) and Brav, Constantinides and Geczy (2002). Here the focus is on the relationship between the time-series properties of consumption and the coefficient of risk aversion. The last set of tests is based on a more structural approach and uses information on the changes in expectations of consumption and asset returns to estimate efficiently the parameters of interest. In all instances, I use the US Consumer Expenditure Survey, and for most of the analysis I maintain the standard assumption of isoelastic utility; however, whenever possible, I check the implications of preferences characterized by habit persistence and of Epstein and Zin-type utility.

The rest of the paper is organized as follows. In Section 2, I review briefly the C-CAPM and derive its testable implications. Section 3 presents the data and some evidence on household financial market participation based on the US Consumer Expenditure Survey. Section 4 discusses the problems that arise when testing the C-CAPM using micro data and allowing for heterogeneous financial market participation. Section 5 presents the results of the tests and Section 6 assesses the evidence and concludes.

2. The Consumption-based Capital Asset Pricing Model

Consider the standard intertemporal optimization problem facing a consumer with access to N different assets. The consumer chooses consumption and portfolio by maximizing expected lifetime utility (appropriately discounted) subject to an intertemporal budget constraint that reflects the intertemporal allocation possibilities. There are three key assumptions behind this model. First, preferences over future random consumption streams are intertemporally additive. Second, individuals can trade assets in a frictionless market. Third, asset markets are complete. When markets are complete, we can construct a "representative" agent, because after trading in complete markets, individuals become homogeneous (at the margin), even though they are initially heterogeneous. Hence, the maximization problem can be written as follows:

$$\max E_t\left[\sum_{s=t}^{T}\beta^{s-t}U(C_s)\right];$$

(1) s.t.

$$\sum_{k=1}^{N}A_s^k=\sum_{k=1}^{N}A_{s-1}^k\left(1+r_s^k\right)+Y_{s-1}-C_{s-1},$$

where C_s denotes the representative individual consumption in period s, Y_s is the consumer's non-asset income, β the discount factor, A^k the amount of wealth held in asset k, whose return is r^k. If asset k is held at t and t+1, the first order condition for the problem is:

(2)
$$U'(C_t)=E_t\left[\beta U'(C_{t+1})\left(1+r_{t+1}^k\right)\right],$$

which can be re-written as:

(3)
$$E_t\left[m_{t+1}\left(1+r_{t+1}^k\right)\right]=1,$$

where $m_{t+1}=\beta\frac{U'(C_{t+1})}{U'(C_t)}$. m_{t+1} denotes the marginal rate of substitution (IMRS) between consumption at time t and at t+1 and is known as the stochastic discount factor. Since (3) must hold for all assets held in non-zero amounts at t and at t+1, it can also be expressed in terms of excess return between, for example, a risky asset k and a risk-free asset f:

(4)
$$E_t\left[m_{t+1}\left(r_{t+1}^k-r_{t+1}^f\right)\right]=0.$$

(3) and (4) impose unambiguous statistical restrictions on the co-movement between consumption changes and asset returns.

A key implication of equation (3) is that equilibrium returns are determined by the intertemporal marginal rate of substitution of consumption. When markets are complete, observed asset returns are sufficient to identify the IMRS. When markets are incomplete, there may be idiosyncratic variation in individual investors' marginal utilities, hence multiple stochastic discount factors that are consistent with (3). However, as Hansen and Jagannathan (1991) show, observed returns can still be used to derive a set of restrictions on the pricing factors. Using only financial data, they restrict the admissible region for the mean and standard deviation of m_{t+1} by constructing minimum variance stochastic variables that vary with returns in the same way as the IMRS in the Euler equation in (3)[6]. The region provides a convenient summary of the sense in which asset market data are anomalous from the vantage point of intertemporal asset pricing theory.

Alternative tests of (3) and (4) can be designed using a more structural approach and the orthogonality conditions that they imply to estimate the preference parameters. A simple test based on the unconditional restrictions that the model in (1) implies can be obtained using the law of iterated expectations and replacing the conditional expectations in (3) and (4) with unconditional ones, so that:

(5) $$E\left[m_{t+1}\left(1+r_{t+1}^{k}\right)\right]=1,$$

(6) $$E\left[m_{t+1}\left(r_{t+1}^{k}-r_{t+1}^{f}\right)\right]=0.$$

Then, given a sufficiently long time series of data on C_t and asset returns, we can calibrate the expectation on the left-hand side of (6) using the sample mean of:

(7) $$e_{t+1}=m_{t+1}\left(r_{t+1}^{k}-r_{t+1}^{f}\right),$$

[6] In practice, Hansen and Jagannathan (1991) focus on unconditional moment restrictions - which can be obtained by applying the law of iterated expectations to the conditional version in (3) -, because they are typically easier to estimate than conditional moments. An extension of part of their analysis based on conditional moments is in Gallant, Hansen and Tauchen (1990).

and determine the preference parameters in m_{t+1} that justify observed behavior. The logic behind this test is that only if $E(e_{t+1}) = 0$ will the agent be at an optimum and unable to improve his welfare by borrowing at the risk-free rate and investing in the risky asset. Note that, as Cochrane and Hansen (1992) emphasize, Hansen and Jagannathan's lower bound on the volatility of m_{t+1} is an even weaker implication of (4) than (6), in that (6) implies the variance bound while the converse is not true.

The relationship upon which Mankiw and Zeldes (1991) rely can be derived from (5) by adding the assumptions of isoelastic utility and of log normal consumption growth and asset returns. Under these hypotheses, the left-hand-side of (5) can be written as:

$$E\left[m_{t+1}(1+r^k_{t+1})\right] = \exp\left\{-\gamma E\left[\log\frac{C_{t+1}}{C_t}\right] + E\left[\log R^k_{t+1}\right] + \delta^k\right\}, \tag{8}$$

where γ is the coefficient of relative risk aversion and the inverse of the elasticity of intertemporal substitution, R^k is the gross return on asset k and:

$$\delta^k = \log\beta + \frac{1}{2}\left(\gamma^2 Var\left(\log\frac{C_{t+1}}{C_t}\right) + Var\left(\log R^k_{t+1}\right)\right) - \gamma Cov\left(\log\frac{C_{t+1}}{C_t}, \log R^k_{t+1}\right). \tag{9}$$

Given (8), taking the log of both sides of (5) yields:

$$E\left[\log R^k_{t+1}\right] - \gamma E\log\frac{C_{t+1}}{C_t} + \delta^k = 0\,. \tag{10}$$

When a risk-free asset is available, (10) and (9) imply

$$E\left[\log R^k_{t+1}\right] - \log R^f_{t+1} = \gamma Cov\left(\log\frac{C_{t+1}}{C_t}, \log R^k_{t+1}\right) - \frac{1}{2}Var\left(\log R^k_{t+1}\right). \tag{11}$$

After some re-arrangement, (11) can be written as:

$$\begin{aligned}\log E\left[R^k_{t+1}\right] - \log R^f_{t+1} &= \gamma Cov\left(\log\frac{C_{t+1}}{C_t}, \log R^k_{t+1}\right) \\ &= \gamma Corr\left(\log\frac{C_{t+1}}{C_t}, \log R^k_{t+1}\right)\cdot\sqrt{Var\left(\log\frac{C_{t+1}}{C_t}\right)Var\left(\log R^k_{t+1}\right)},\end{aligned} \tag{12}$$

which is the unconditional version of (4) that Mankiw and Zeldes (1991) use to estimate risk aversion. Since the coefficient of relative risk aversion, γ, is the only parameter to be determined, it can be identified by computing the sample equivalents of the unconditional population moments in (12). This approach is appealing because it relates the preference parameters of interest directly to the time-series properties of consumption growth and asset returns. However, since it relies on a single orthogonality condition to estimate γ, it tends to be rather inefficient. Hence, the usefulness of a conditional version of the first order condition to the maximization problem in order to improve the precision of the estimates by adding information about predictable movements in expected consumption growth rates and asset returns.

Following Hansen and Singleton (1983), we can log-linearize the (conditional) Euler equation in (3) under the assumption of isoelastic preferences and obtain:

(13) $$\log \frac{C_{t+1}}{C_t} = \lambda_t^k + \gamma^{-1} \log R_{t+1}^k + \varepsilon_{t+1}^k.$$

Under the assumption of jointly log-normal consumption growth and asset returns, $\lambda^k{}_t$ is given by:

(14) $$\lambda_t^k = \frac{1}{\gamma}\log\beta + \frac{1}{2}\left(\gamma Var_t\left(\log\frac{C_{t+1}}{C_t}\right) + \frac{1}{\gamma}Var_t\left(\log R_{t+1}^k\right)\right) - Cov_t\left(\log\frac{C_{t+1}}{C_t}, \log R_{t+1}^k\right),$$

where $\mathrm{Var}_t(.)$ and $\mathrm{Cov}_t(.)$ denote variance and covariance conditional upon the information available at time t. The residual $\varepsilon^k{}_{t+1}$ includes expectation errors and is given by:

(15) $$\varepsilon_{t+1}^k = \log\left(\frac{C_{t+1}}{C_t}\right) - E_t\left[\log\left(\frac{C_{t+1}}{C_t}\right)\right] + \frac{1}{\gamma}\left(\log R_{t,t+1}^k - E_t\left[\log R_{t,t+1}^k\right]\right).$$

If there are instruments that are uncorrelated with $\varepsilon^k{}_{t+1}$ and with the innovations to $\lambda^k{}_t$, (13) can be estimated using the generalized method of moments. Furthermore, if the model is over-identified, it is possible to test for the predictability of deviations from the Euler equation. Notice that we could have estimated risk aversion applying the method of moments directly to the non-linear Euler equation in (3). However, Monte Carlo studies by Vissing-Jørgensen (1999) show that in the presence of measurement error, which is common in micro data, the estimator based on non-linear Euler equation has poorer properties than estimators based on the log-linearized version in (13).

3. Who Owns Financial Assets in the US?

When deciding how to structure their portfolio, US households face an array of options, ranging from simple bank accounts to highly sophisticated real estate investment trusts. Nevertheless, the portfolios of American households tend to be very simple and safe and characterized by a high degree of heterogeneity. Most households hold a bank account, but relatively few have an investment in stocks and, overall, the ownership of many types of assets is heavily concentrated at the top of the income distribution. Hence, the idea of testing whether the inclusion of non-shareholders consumption in the pricing factor for shares is at the root of the empirical inconsistency of the C-CAPM.

This section describes the extent and patterns of asset ownership in the US and the degree of heterogeneity in financial market participation using the US Consumer Expenditure Survey (CEX), the data set that I employ for assessing the empirical performance of the C-CAPM. Although the information on wealth and savings is less accurate than in the Survey of Consumer Finances (SCF), with financial wealth disaggregated into only four categories, the CEX has two main advantages over the SCF. First, it provides 17 years of data, available on a consistent basis at an annual frequency, against 5 waves of the SCF. Second and more importantly, it contains detailed information on consumption.

The CEX is a representative sample of the US population. It is a rotating panel in which interviews occur continuously throughout the year, each consumer unit being interviewed every three months over a twelve-month period, apart from attrition. The sample I use consists of 34,033 households[7] and covers the period from 1982 to 1996, first quarter. Each quarterly interview collects household monthly expenditure data on a variety of goods and services for the previous three months. For my analysis, however, I use only the

[7] See the Appendix for details on the data and on the criteria of household selection and exclusion in the sample used for the analysis.

information for the month preceding that of interview,[8] and the consumption measure I use is real seasonally adjusted household expenditure[9] on non-durable goods and services.

The survey contains detailed information on demographic and socio-economic characteristics of the household and its members. During the last interview households fill in a financial supplement and provide information on their holdings of four types of assets: 1. checking, brokerage and other accounts; 2. savings accounts; 3. US savings bonds; 4. stocks, bonds, mutual funds and other securities. “US savings bonds” and “stocks, bonds, mutual funds and other securities” are added together and labeled “risky assets” and the households that have non-null holdings are defined as “shareholders”. Hence, shareholders will inevitably include many households that do not hold stocks, but only bonds. However, this “imperfect” distinction involves a bias against finding significant differences between shareholders and non-shareholders. As a consequence, it will bias my analysis against, rather than in favor of my prior that the C-CAPM is an accurate description of shareholders’ behavior but should not be expected to hold for non-shareholders. Overall, in my sample, 77%, 68% and 32% of households hold checking accounts, savings accounts and risky assets, respectively.

Tables 1 and 2 report the patterns of asset ownership in the US, according to the CEX. Limited participation is particularly pronounced for the risky asset market and can hardly be explained within the framework of traditional portfolio choice models, which imply widespread ownership of stocks given their high return. Table 1 gives the percentages of households with the possible combinations of assets available. I distinguish among households with no assets (15% of the sample), with only checking accounts (14%), only savings accounts (6%), only checking and savings accounts (33%), and both risky assets plus

[8] In practice, many households report to spend exactly the same amount in each of the three months preceding that of interview. This is likely to be a consequence of the fact that this information is based on recall (and not, for example, on diaries) and households are dividing by three their quarterly expenditure. Thus, using all three observations would not add much information.

[9] Nominal consumption is deflated by means of household-specific indices based on the Consumer Price Index provided by the Bureau of Labor Statistics. The individual indices are determined as geometric averages of elementary regional price indices, weighted by the shares of household expenditure on individual goods. See Attanasio and Weber (1995) for a more extensive discussion of these indices. The seasonal adjustment uses twelve month dummies.

checking and/or savings accounts (32%).[10] Two features are particularly remarkable: first, many households do not hold any financial asset; second, most hold just few of the available assets. The most significant trend is the increase in the share of households with risky assets, from 28% in the early 1980s to 34% in the 1990s, which is consistent with the evidence based on the SCF.[11]

Table 2 relates portfolio composition to some household characteristics, such as age, education, gender, race, household composition, income and expenditure, pooling all years together. With respect to Table 1, I have grouped together the households holding just checking or savings accounts (riskless asset holders), who represent around 53% of the sample. Hence, the Table distinguishes between households without assets, households with just riskless assets and households holding also stocks and bonds. Financial market participation, and especially risky asset ownership, appears to increase with age, education and income, which is consistent with the theories that view financial market participation as costly. The observed correlation between income and education, on the one hand, and financial market participation, on the other, suggest that such costs may be monetary or may be related to all the opportunity costs of investors' time in gathering and processing the information needed to participate to financial markets (see Haliassos and Bertaut (1995) for further discussion). Also, the households participating in financial markets are more likely to be white and headed by a male. Risky asset holding is quite uniformly distributed around the country, whereas no-asset holders are concentrated in the South. In terms of size and composition, households with risky assets are somewhat larger and more likely to have children than those holding just checking and savings accounts, but are smaller and less likely to have children than those holding no assets. The latter are often female-headed, and a larger share are non-white. Table 2 reports also median household non-durable consumption and mean and median quarterly consumption growth, which, as expected, are lowest among those households with no financial assets and highest among those holding risky assets. The last row reports the rate of growth of consumption of the representative agent, i.e. the household consuming mean consumption. The quarterly growth rate for the

[10] 147 households that reported holding only risky assets are ignored in these tables.

[11] See Poterba and Samwick (1995) and Bertaut and Starr-McCluer (1999).

representative holder of risky as well as riskless assets is marginally higher than for the representative riskless asset holder, and substantially higher than for the representative household with no financial assets.

The choice of holding risky assets is investigated further in Table 3, which reports the results of the estimation of a bivariate choice (probit) model for risky asset ownership. The individual probabilities of holding risky assets are determined as a function of a polynomial in the age of the household head, of his education, gender, race, marital status, earnings, of the presence of children in the household and of the area of residence. A set of year dummies is also included. The likelihood of holding financial assets appears to be increasing, although not linearly, in age and also in education. The dummy for children has a positive sign, but is hardly significant. Single-person, female-headed and non-white households appear to be less likely to hold shares. Finally, the probability of holding shares is highest in the North-East and Mid-West and is strongly increasing in earnings.

Table 4 and Figure 1 summarize the evidence on the real quarterly returns that I use in the analysis. As risky return, I take the return on the Standard & Poor's 500 Composite Index (S&P500 CI). Unsurprisingly, over the period 1982-1996, on average, the risky return is substantially higher and more volatile than the return on 3-month Treasury Bills (T-bills), which I take as riskless asset. The mean quarterly premium is more than 2.5 %.

4. Accounting for Household Portfolio Heterogeneity

The inference based on tests of the C-CAPM that use aggregate data would be valid only if non-shareholders' expenditure co-varied with share returns in the same way as shareholders' spending. Otherwise, individual level data providing information on consumption and financial wealth are needed to distinguish between those for whom the Euler equations in (3) can be expected to hold and those for whom they cannot.

The distinction between asset holders and non-holders can be carried out in two ways. One is to focus on actual asset holdings, as in Mankiw and Zeldes (1991), who classify as shareholders those who hold an amount of shares above a certain threshold. [12] Yet, this way of proceeding is likely to bias the analysis of the link between consumption and asset returns conditional upon participation, due to the endogeneity of asset holding with respect to expenditure. In fact, a shock to consumption can induce the household to invest in shares, while it had not before, or to liquidate its assets. Further, suppose that the instantaneous utility function of household h is given by:

$$U\left(c_{h,t}, v_{h,t}\right) = (1-\gamma)^{-1} c_{h,t}{}^{1-\gamma} \exp\left(v_{h,t}\right), \tag{16}$$

where $c_{h,t}$ denotes household expenditure and $v_{h,t}$ captures individual unobserved heterogeneity. If $v_{h,t}$ is correlated with the decision to hold shares, grouping households on the basis of actual ownership, can yield inconsistent estimates.

An alternative procedure has been suggested by Attanasio, Banks and Tanner (2002), who distinguish between shareholders and non-shareholders on the basis of the individual likelihood of holding shares. This approach can be thought of as an instrumental variable procedure, where, in the first stage, participation probabilities are determined and, in the second, the estimated probabilities are used to instrument actual participation and to study its effects on consumption.

I adopt the latter approach and compute the individual likelihood of shareholding using the results from the probit estimation of the probability of financial market participation reported in Section 3. More specifically, I label as “predicted shareholders” those households whose estimated probability of holding risky assets at time t[13] is above a certain threshold. In order to account for changes in participation over time, I allow for a variable threshold. The cut-off point is set equal to the probability that ensures that the portion of predicted shareholders is the same as that of households that actually hold stocks and bonds in each

[12] Mankiw and Zeldes (1991) adopt three definitions of shareholders: those reporting any amount of shares, those with at least $1,000, and those with $10,000 or more.

[13] In the CEX, households report their asset holdings at the time of the last interview, i.e. just before leaving the survey, which can be labeled as time t+1. For the analysis based on the Euler equation for consumption at t and t+1, what matters is participation status at time t. Once the probit for the probability of participating at t+1 is estimated, the probability of participating at time t can be easily recovered using the estimated coefficients since the model is specified in terms of variables that are fully observable both at t+1 and at t (e.g. earnings) or evolve predictably over time (e.g. age).

year covered by the sample.[14] Then, for each couple of adjacent dates, once the group of likely shareholders and non-shareholders is defined on the basis of their probability of participating in the first of the two periods, I can compute consumption growth for the individuals of each group or for the representative agent of the group.

The distinction based on individual participation probabilities is such that predicted shareholders end up including households that in fact do not hold shares and excluding some that do. This error could be avoided only if the participation model predicted individual choices exactly, but this is unrealistic because of unobservable heterogeneity, which is the source of bias also in an analysis based on actual asset holdings because of the correlation between participation and unobservable preference parameters. However, when participation is determined exogenously, the individual unobservable heterogeneity that is left in the error has zero mean in the groups of predicted shareholders and non-shareholders, which ensures the consistency of the results. The main implication of this error concerns the choice of the instruments needed to estimate the parameters of interest, using (13), which must be uncorrelated with this unobservable term. Instruments dated t-1 and earlier address this issue.

Given asset holding status at time t, I can look at the time series properties of consumption growth of predicted shareholders and non-shareholders either by taking into account the cross-sectional heterogeneity in consumption growth rates or by assuming a representative agent (within each group). The former option amounts to computing individual level consumption changes using the panel dimension of the data set, which allows us to construct up to three quarterly consumption growth observations for each household. The latter option amounts to using the growth rate of average consumption. In this instance, since 20% of the sample is replaced each quarter, care should be taken to ensure that households are matched across quarters. With matching, the (quarterly) growth rate of average consumption is given by: $\frac{C_{t+3}}{C_t} = \frac{H_t^{-1}\sum_{h\in H_t} c_{h,t+3}}{H_t^{-1}\sum_{h\in H_t} c_{h,t}}$. C_{t+3} is computed using

[14] The fraction of shareholders varies across years between 27 and 35%. The results are robust to alternative thresholds.

only the observations on those H_t households that were also interviewed during month t.[15] With matching, the growth rate of consumption is not affected by systematic differences in the consumption levels of households entering the survey at different dates. Otherwise, if the expenditure of those that have just entered differs systematically from that of the households who are completing their participation, an additional element of noise, and potential source of bias, would arise. Further, matching addresses the issue of composition effects due to changing probabilities of ownership and to changing probability thresholds that can cause the composition of the group of predicted shareholders to change over time. Systematic changes in the group of predicted shareholders might create serious distortions in the presence of unobserved heterogeneity because the IMRS would encompass both genuine changes in consumption and composition effects. To make this point clear, notice that the IMRS corresponding to the isoelastic utility function specified in (16), with multiplicative heterogeneity, is given by: $IMRS_{t,t+3} = \left(c_{h,t+3}\right)^{-\gamma} c_{h,t}{}^{\gamma} \exp\left(v_{h,t+3} - v_{h,t}\right)$. If the cross-sectional mean of $v_{h,t}$, $E\left[v_{h,t}\right]$ for $h \in H_t$, is constant and households are matched forward, unobserved heterogeneity does not cause problems, as the terms in v drop out.

The choice of aggregating consumption growth cross-sectionally or assuming a representative agent should account for the presence of measurement error, which can be a serious issue for any consumption study based on micro data. Under the standard assumptions that measurement error is multiplicative[16], uncorrelated across households and independent from the true consumption level and asset returns, the Euler equations in (3) yield consistent estimates of the parameters characterizing the curvature of the utility function, but not of the discount factor. This is the case whether the Euler equations are log-linearized or not and whether there is a panel for the estimation or consumers are aggregated

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[15] Without matching, it would be: $\frac{C_{t+3}}{C_t} = \frac{H_{t+3}^{-1} \sum_{h \in H_{t+3}} c_{h,t+3}}{H_t^{-1} \sum_{h \in H_t} c_{h,t}}$, with H_{t+3} including both households that have entered the survey at an earlier date and households that are new to the survey.

[16] Assuming multiplicative measurement error is reasonable since it is more likely for agents to misreport their consumption by a fraction of the actual figure than by an amount that is constant and independent of actual expenditure. Under this hypothesis, observed consumption $c_{h,t}$ would correspond to $c_{h,t} = c_{h,t}^{*} \cdot \zeta_{h,t}$, with $c_{h,t}^{*}$ denoting actual consumption and $\zeta_{h,t}$ reflecting measurement error.

within each period. Nevertheless, when the standard deviation of the measurement error is large, in practice, log-linearized Euler equations are preferable and give estimates with better properties. For Hansen and Jagannathan bound analysis of stochastic discount factors, when consumption is measured with error, both the mean and the standard deviation of the pricing factor (the IMRS) are biased, unless we assume a representative agent. Within a representative agent framework, measurement error has no effects as *H*®¥, if its distribution is stationary over time.[17] For this reason, whenever dealing with non-linear functions of the parameters of interest, I make the representative agent assumption.

Another reason for using consumption averages is to reduce the impact of the observations associated with falling expenditure. The IMRS involves consumption ratios raised to $-\gamma$. If household expenditure falls over two adjacent dates, the consumption ratio will be less than one and the contribution to the IMRS will be greater, the more risk averse the agent is. This implies that for high values of risk aversion those observations associated with negative consumption growth prevail and the IMRS sample moments end up being crucially determined by these observations and turn out to be poor estimates of the population moments. Using consumption averages helps deal with this problem.

Nevertheless, the representative agent assumption is not innocuous when focusing on non-linear quantities such as the IMRS. In fact, using ratios of means instead of means of ratios can create a source of bias. However, if the cross-sectional variance of consumption growth is uncorrelated with returns, the representative agent assumption yields the same results as when taking cross-sectional heterogeneity into account.[18] Another, alternative assumption that would guarantee no aggregation bias is that the rate of growth of individual consumption is uncorrelated with initial consumption and consumption is log-normally distributed in the cross-section. When the IMRS is made a linear function of parameters, as in log-linearized Euler equations, the representative agent assumption is not a problem.

Before turning to the empirical evidence, it is worth mentioning that the analysis relies on quarterly consumption growth rates at monthly frequencies, which implies overlapping

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[17] See Vissing-Jørgensen (1999) for an extensive discussion of these issues.

[18] See Constantinides and Duffie (1996) for a discussion of these issues.

observations. In addition, households are interviewed quarterly and contribute up to 4 observations on expenditure. All this affects the structure of the residual of the log-linearized Euler equations, which results from two components: an expectation error and a measurement error. The former is white noise, or at most an MA(1) due to presence of time aggregation effects, at the household level. Nevertheless, because of the overlapping of the observations on consumption growth, the individual expectation components in ε_{t+3}[19] can be expected to be correlated with the expectation components in ε_{t+2} and ε_{t+1}, yielding an MA(2). The measurement error generates an MA(6) component and is due to the fact that, in the absence of attrition, the household that enters the survey at t contributes to the IMRS dated t+3, t+6 and t+9. Hence, overall, the resulting residual of (13) is an MA(6).

5. Empirical Implications of Portfolio Heterogeneity

This section provides extensive evidence on the hypothesis that the C-CAPM is a satisfactory description of the equilibrium relationship between consumption and asset returns once we price assets on the basis of the consumption growth of shareholders only. Three types of tests are carried out, each emphasizing different features of the data. The first aims at verifying the implications of portfolio heterogeneity for the evidence from Hansen and Jagannathan bound-based analysis. This is a diagnostic test that focuses mainly on the role of the IMRS volatility for the C-CAPM. Next, I calibrate the C-CAPM to assess the extent of the differences in the consumption series of shareholders and non-shareholders and their impact on the estimation of preference parameters based on the unconditional restrictions that the model implies. Finally, I estimate the parameters of interest exploiting the information about time variation in expected consumption growth and asset returns, using an instrumental variables approach that allows to test for the overall suitability of the C-CAPM in describing the data.

[19] The left-hand side variable of the equation to be estimated is: $\sum_{h\in H_t} \log \frac{c_{h,t+3}}{c_{h,t}}, \quad t=1,T-3$, with t denoting the month.

The model is estimated/tested using information on the subset of shareholders, with the grouping based on the estimated individual probabilities of holding shares, as described earlier. As a benchmark I take the representative agent of the CEX sample used in the analysis, whose expenditure is known not to match US aggregate per-capita expenditure exactly, especially in the levels.[20] As a consequence, comparing the evidence on CEX shareholders with that on the "US aggregate" representative agent would be problematic, since any differences might be due not just to a proper accounting of household heterogeneity, but also to fundamental differences in the data sets. Hence, I compare the evidence on CEX shareholders with that on an agent who consumes mean CEX expenditure. The Tables also report the results of the estimation/testing of the model based on the data for predicted non-shareholders to assess the suitability of the theory to explain their behavior.

5.1 Hansen and Jagannathan Bounds

Hansen and Jagannathan (1991)'s framework of analysis consists in a lower bound on the volatility of the stochastic discount factors that would be consistent with a given set of asset return data. This approach is non-parametric and applies to a rich class of pricing models. It can be thought of as complementary to those based on more traditional tests of overidentifying restrictions, such as Hansen and Singleton (1982, 1983), whose success is mixed since they do not provide any indication as to how the model should be changed in order to correct or improve its specification. In addition, very often, such tests focus excessively on the goodness of the specification and ignore the predictive potential.

Figures 2 through 4 show the implications of the distinction between shareholders and non-shareholders. In the figures, Hansen and Jagannathan volatility bounds coincide with the solid line and have been determined using the quarterly returns on the S&P500 CI and on US T-bills over the period 1982:1-1996:1. The sequences of crosses and triangles represent the mean (on the horizontal axis) and standard deviation (on the vertical axis) of the IMRS of predicted shareholders and non-shareholders, respectively, for different values of risk aversion. The dashed lines refer to the CEX representative agent. The coefficient of risk

[20] See Sabelhaus (1998) for a discussion of the measurement problems affecting the CEX.

aversion starts from 0.5 and increases by steps of 0.5. The quarterly discount rate is set to 2 %. Each figure incorporates alternative hypotheses regarding the structure of preferences. Figure 2 is based on the (standard) assumption of isoelastic utility, which implies that the IMRS is defined as $\beta\left(\bar{c}^{g}_{t+3}/\bar{c}^{g}_{t}\right)^{-\gamma}, g = s, ns$. g denotes the groups of predicted shareholders (*s*) and non-shareholders (*ns*). $\bar{c}^{g}_{t}$ is (group) mean expenditure on non-durables and services; g is the Arrow-Pratt coefficient of relative risk aversion. Figure 3 is based on the hypothesis that preferences exhibit habit persistence, which implies an IMRS defined as $\frac{\beta\left(\bar{c}^{g}_{t+3}-\delta\bar{c}^{g}_{t}\right)^{-\gamma}-\beta^{2}\left(\bar{c}^{g}_{t+6}-\delta\bar{c}^{g}_{t+3}\right)^{-\gamma}\delta}{\left(\bar{c}^{g}_{t}-\delta\bar{c}^{g}_{t-3}\right)^{-\gamma}-\beta\left(\bar{c}^{g}_{t+3}-\delta\bar{c}^{g}_{t}\right)^{-\gamma}\delta}$, where d measures the impact of past consumption on current utility. To avoid composition effects, in this instance, the IMRS is computed consistently using the sample of households participating to all four interviews. d is set equal to 0.15 and the choice is dictated primarily by empirical considerations to prevent the IMRS from turning negative. To my knowledge, no readily available estimate of d is available and, in general, time-non-separable preferences have not had much success in empirical applications because non-negligible degrees of habit are problematic to use empirically, since the corresponding IMRS are often negative and therefore lack economic meaning, especially within micro data-based frameworks. Finally, in Figure 4, I consider the case of Epstein and Zin (1989, 1991) utility. The IMRS is defined as $\left[\beta\left(\bar{c}^{g}_{t+3}/\bar{c}^{g}_{t}\right)^{-\rho}\right]^{\frac{1-\gamma}{1-\rho}} R_{t+3}^{\frac{1-\gamma}{1-\rho}-1}$. R_{t+3} is the quarterly return on the market portfolio, which, in the case considered, corresponds to the return on the S&P500 CI.[21] r is the inverse of the intertemporal elasticity of substitution and is set equal to 5 and 0.5, implying an elasticity of 0.2 and 2, respectively. In the literature, there is no consensus as to the value of the elasticity of substitution. Hall (1988) supports the view of an elasticity below one. Other papers, such as Mankiw *et al*. (1985), obtain higher values that are generally above one. A relatively high elasticity of substitution would help solve the risk-free rate puzzle, concerning the empirical inconsistency of consumption growth with the risk-free return. Since the former is found to be rather high, whereas the latter is quite low, for the two to be reconciled the elasticity of substitution must be much greater than the inverse of the risk-aversion coefficient that is consistent with the observed

[21] See Epstein and Zin (1989, 1991) for a thorough discussion of this class of preferences.

risk premium. Kocherlakota (1996) shows that if the link between the two parameters is broken, when γ ranges from 2 to 10 the elasticity of substitution that reconciles consumption growth and the risk-free returns varies between 1 and 10.

Figure 2 clearly shows the importance of distinguishing between shareholders (crosses) and non-shareholders (triangles) and the implications of the representative agent assumption (dashes), when preferences are isoelastic. The IMRS based on per-capita expenditure is consistent with observed returns for a risk aversion around 9. Shareholders' sequence never crosses the bound, but it comes close to it for a much lower level of risk aversion, around 3. Instead, non-shareholders' risk aversion sequence is consistent with the bound for a higher γ, above 11. Notice that the picture based on CEX data differs considerably from that of Hansen and Jagannathan (1991) based on US aggregate data, which imply consistency of the IMRS with the bounds for a risk aversion above 200. The differences in the sequences for the CEX representative agent and for the US representative agent appear to be due to the fact that idiosyncratic components make the series based on individual data much more volatile than its aggregate counterpart. Idiosyncratic risk *vis à vis* the relatively small number of shareholders versus that of non-shareholders might also partly explain the sharp increase of predicted shareholders' IMRS mean as the coefficient of risk aversion increases. Still, overall, Figure 2 provides a first piece of evidence that the C-CAPM may be an acceptable description of households' behavior, i.e. may be consistent with the data for low risk aversion, when portfolio heterogeneity is properly accounted for and non-participants in financial markets – whose expenditure cannot be expected to satisfy the restrictions – are excluded.

Figure 3 displays the IMRS series in the case of preferences characterized by habit persistence. As expected, the complementarity of consumption over time increases the IMRS volatility with respect to the case of isoelastic preferences, which has two effects. On the one hand, at very low levels of risk aversion the IMRS sequences are relatively closer to the bound. On the other, as γ increases, the mean and standard deviation of the IMRS increase at a rate that is too high for the series to ever cross the bound. However, as with isoelastic utility, the IMRS sequence of predicted shareholders (crosses) is closer to the Hansen and Jagannathan region than the sequence of the representative agent (dashes) and much more so than the sequence of non-shareholders (triangles). As the degree of habit persistence increases (not shown), so does the variance of the IMRS. Yet, for degrees of habit above 0.2, the IMRS is often negative and the mean-standard deviation series become extremely erratic.[22] This is symptomatic of the problems arising with this type of utility, which is incompatible with sudden changes in consumption, which are quite frequent in individual level consumption series.

Figure 4 considers the implications of Epstein and Zin utility for different values of the elasticity of substitution. The sequences of $\times$ and ∇ refer to an elasticity of 0.2; those of + and Δ refer to an elasticity of 2. When the elasticity of substitution is less than 1, the IMRS series of predicted shareholders ($\times$) again turns out to be closer to the bound than those of the representative agent (--) or of non-shareholders (∇). With an elasticity of substitution greater than 1, the IMRS for predicted shareholders (+) is fully consistent with the bound for a risk aversion between 3 and 3.5. Interestingly, the IMRSs for the representative agent (- -) and for predicted non-shareholders (Δ) are also consistent with the bound for similar levels of risk aversion. This result is at odds with the hypothesis that accounting for portfolio heterogeneity helps to explain the empirical inconsistency of the C-CAPM. Yet, it does provide evidence favorable to the theories that assign a role to preference non-separabilities.

Overall, the analysis carried out in this section implies that if preferences are isoelastic, then for realistic values of risk aversion relaxing the assumption of representative agent and focusing on shareholders substantially reduces the distance of the IMRS-based pricing factor from the Hansen and Jagannathan bound. The comparison between alternative utility specifications yields mixed results. On the one hand, it is clear that preference non-separabilities as induced by habit persistence can reduce the distance from the bound by increasing the volatility of the IMRS. However, the empirical tractability of this type of preferences (and therefore the possibility of testing) is limited due to the "constraints" that habit persistence sets on the rate of growth of consumption. On the other hand, the evidence regarding the role of heterogeneity in a framework that relaxes expected utility is inconclusive and needs further analysis. It is important to remember that the results based on Epstein and Zin utility depend on the assumption that the return on the S&P500 CI is a good

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[22] δ_s smaller than 0.15 yield results that are similar to the ones reported here.

proxy for the return on the market portfolio, i.e. that it counts most of the assets that make up household wealth and that the factors excluded are negligible. The validity of this assumption should be studied further.

5.2 Calibration of the Model

In this section I calibrate the C-CAPM and analyze the implications of portfolio heterogeneity for the non-conditional restrictions it implies. The calibration of the model is complementary to the analysis based on the Hansen and Jagannathan bounds. The latter focuses on the role of the IMRS volatility in the equity premium puzzle; while the former looks directly at the relationship giving rise to the puzzle and to the role of the correlation between consumption and returns.

I proceed by estimating the model both in its non-linear form - which allows us to examine the implications of alternative specifications of the felicity function, but is not as robust to measurement error - and after log-linearizing it, in the spirit of Mankiw and Zeldes (1991).

5.2.1 The non-linear model

Table 5 shows the results of risk aversion estimation based on the unconditional version of the Euler equation reported in (6), under the hypotheses of isoelastic preferences, habit persistence and Epstein and Zin utility. The Table reports the minimum value of the risk aversion parameter that is consistent with observed behavior, i.e. that makes $\bar{e}^{g} = (T-1)^{-1}\sum_t e^{g}_{t+3}$, $g = s, ns$, statistically indistinguishable from zero, where $\bar{e}^{g}_{t+3} = \beta \frac{U'(\bar{c}^{g}_{t+3})}{U'(\bar{c}^{g}_{t})}\left(R^{S\&P500}_{t+3} - R^{Tbill}_{t+3}\right)$. As before, g denotes the groups of predicted shareholders (s) and non-shareholders (ns). $U'(.)$ is the marginal utility of consumption. $\bar{c}^{g}_{t}$ is mean (group) expenditure on non-durables and services. Standard errors are in parentheses.[23]

[23] The standard errors have been corrected to account for the MA(6) structure of the residual. The problem of non-positive definite variance covariance matrix in finite samples has been taken care of by using a set of weights, as in Newey and West (1987).

The Table shows that when preferences are isoelastic accounting for household heterogeneity lowers the coefficient of relative risk aversion that reconciles observed expenditure behavior with asset returns. In fact, it drops from 9 for the whole CEX sample to 7 for predicted shareholders, against 13.5 for non-shareholders. The evidence does not change in any important respect when allowing for complementarity in consumption over time. Shareholders' expenditure choices are consistent with the model if their risk aversion is around 7, compared with 10.5 for the representative agent. Non-shareholders' risk aversion appears to be somewhat higher, but ε^{ns} is negative. A negative ε^{ns} is inconsistent with the maximization model, but here it is likely to result from the fact that empirically, in the presence of habit formation there is nothing to prevent the instantaneous utility from being negative.

The evidence regarding Epstein and Zin utility is consistent with that emerging from the Hansen and Jagannathan-type analysis in Figure 4. When the elasticity of substitution is set to 2, predicted shareholders are largely indistinguishable from non-shareholders and the C-CAMP turns out to be consistent with the empirical evidence for a relative risk aversion around 8, with no need to account for portfolio heterogeneity. When the elasticity of substitution is less than 1, the estimated risk aversion for predicted shareholders is lower than for the representative agent or for non-shareholders but still turns out to be above 60.

5.2.2 The log-linear model

Table 6 shows the risk aversion estimates based on the calibration of the unconditional log-linearized C-CAPM of equation (12), whose sample analogue implies:

$$(17)\quad \gamma^g = \frac{\log \overline{R^{S\&P500}} - \log \overline{R^{T-bill}}}{\hat{Corr}\left(\overline{\Delta \log c^g_{t+3}}, \log R^{S\&P500}_{t+3}\right) \cdot \sqrt{\hat{Var}\left(\overline{\Delta \log c^g_{t+3}}\right) \hat{Var}\left(\log R^{S\&P500}_{t+3}\right)}}, \qquad g = s, ns,$$

where $\overline{R^i} = \frac{1}{T-3}\sum_{t=1}^{T-3} R^i_{t+3}$, $\quad i = S\&P500,\ T\text{-}bills$; $\hat{Corr}(.)$ and $\hat{Var}(.)$ denote the sample correlation and variance; $\overline{\Delta \log c^g_{t+3}} = \frac{1}{H^g_t}\sum_{h \in H^g_t} \Delta \log c^g_{h,t+3}$ with H^g_t denoting the set of households interviewed both at date t and at t+3. The rate of growth of consumption of predicted shareholders (g=s) and non-shareholders (g=ns) is based on the geometric (group) mean of individual consumption ratios, which reduces the impact of measurement error and also of outliers. Hence, I use individual level rates, which were unsuitable when the variable examined was the (non-linear) IMRS.

The Table reports the correlation between consumption growth and the (log) return on the S&P500 CI, the standard deviation of consumption growth and the corresponding relative risk aversion coefficient. When the representative agent assumption is adopted, the implied relative risk aversion coefficient is 60.[24] Allowing for portfolio heterogeneity and singling out shareholders, it drops to 44.[25] The difference stems from differences in the pattern of consumption *vis à vis* asset returns between shareholders and non-shareholders. In fact, as expected, shareholders' consumption growth is more volatile than non-shareholders' and the correlation with risky asset returns is stronger.[26] Hence, since non-shareholders' expenditure does not co-vary with returns (as opposed to shareholders'), including their expenditure in the consumption measure used to price shares will bias the inference. Since the estimates are the result of a back-of-the-envelope calculation, standard errors are not reported, but as there is only one moment restriction, standard errors are probably quite large. Overall, the results based on the calibration of the model are qualitatively similar to those obtained within Hansen and Jagannathan's framework: shareholders' expenditure is more volatile and correlates more with the return on shares than non-shareholders and this helps explain why the risk aversion coefficient that reconciles their consumption profile with observed return is substantially lower than that of the representative agent.

By comparison with Mankiw and Zeldes (1991), the results obtained in this section differ quantitatively, but not qualitatively. In fact, Mankiw and Zeldes find somewhat larger differences in the time-series properties of consumption of shareholders and non-

[24] The coefficient of risk aversion based on the geometric mean of the individual consumption ratios for the entire sample is 67.

[25] Calibrating the model using arithmetic means, i.e. under the assumption of (intra group) representative agent, slightly reduces the differences between predicted shareholders and non-shareholders: for the former the estimated relative risk aversion is 48, for the latter 83 (instead of 92).

[26] The cells used to compute average consumption are not very large, and they differ in size between shareholders and non-shareholders. Hence, sampling error can induce part of the time series variability of consumption growth. If we control for sampling error in the variance due to cell size, the relative risk aversion coefficients for shareholders and non-shareholders become 52 and 118, respectively.

shareholders, but estimate the coefficient of risk aversion for the whole PSID sample, for shareholders and for non-shareholders at 100, 35 and 262, respectively.

5.3 Estimation of the Model

The results of instrumental variables estimation of the log-linearized C-CAPM in (13) are reported in Table 7. The relation estimated is the following:

$$\overline{\Delta \log c_{t+3}^{g}} = \lambda^{i} + \gamma^{-1} \log R_{t+3}^{i} + \varepsilon_{t+3}^{g,i} \qquad g = s, ns \quad i = S\&P500, T\text{-}bills\,. \tag{18}$$

The left-hand-side variable of (18) is the rate of growth of consumption of predicted shareholders (g=s) and non-shareholders (g=ns), computed as geometric (group) mean of individual consumption ratios. (18) is estimated for stock returns and for T-bills separately at first and then jointly to improve the efficiency by exploiting the cross-equation correlation of the residuals arising from the correlation of the expectational errors. The evidence is compared with that obtained using as left-hand-side variable the rate of growth of the expenditure of the sample representative agent.

The estimation of the parameter of interest γ is based on the linear generalized method of moments (GMM) and relies on the orthogonality of $\varepsilon_{t+3}^{g,i}$ to the information on the changes over time of households' expectations on consumption and returns. Hence, it is more efficient than the estimation based on the calibration of the model, which relies on a single restriction. The choice and timing of the instruments must account for the MA(6) structure of the residual in (18), for the endogeneity of asset returns with respect to the expectational components of the error, as given in (15), and for the quarterly frequency of the left-hand side variable. Moreover, if the variance and covariance terms in (14) are not constant, their stochastic components enter the residual, which does not cause problems as long as they are uncorrelated with the instruments. Hence, as instruments, I use the quarterly rate of growth of consumption (lagged three and four quarters), the quarterly returns on the S&P500 CI and on T-bills (lagged one to four quarters), the rate of inflation (lagged one to four quarters) and the risk and term spreads (lagged one quarter). The latter are computed as differences between the yield on an index of BAA-rated corporate bonds and the yield on an index of AAA-rated corporate bonds (risk spread) and between the yield on an index of long-term US government bonds and the yield on three-month T-bills (term spread).[27] Lagged stock and bond returns are used in many previous studies as instruments. The use of risk and bond horizon premiums is motivated by the findings of Fama and French (1989) that they have strong predictive power for share returns. For evidence on the predictability of asset returns using the other instruments, see, for example, Attanasio and Weber (1995). First-stage regressions are at the basis of the parsimonious list of over-identifying instruments that I have chosen to use.[28]

The first set of rows of Table 7 display the results when using the S&P500 CI as right-hand-side variable. Under the assumption of isoelastic utility, the point estimate of γ^{-1} implies a relative risk aversion coefficient of 8.3 for predicted shareholders, whereas for non-shareholders γ^{-1} is negative, which is symptomatic of a violation of the Euler equation. For the representative agent it is positive but undistinguishable from zero. When using the return on T-bills (second set of rows), for predicted shareholders, the point estimate implies a relative risk aversion coefficient of around 1.4. For predicted non-shareholders and for the representative agent, it is around 7.7 and 5.2, respectively, although the estimates of γ^{-1} are not statistically significant. In the last set of rows, the Table displays the results from the joint estimation of the equations for the return on the two assets, which imposes an identical value for γ^{-1}. For predicted shareholders, the point estimate implies a risk aversion coefficient around 7.5; for the other two groups it is hardly significant, and for non-shareholders it is negative.[29] The interpretation of the constant terms is not straightforward as they include the discount factor and the unconditional second moments of asset returns and

[27] The standard errors have been corrected to account for the MA(6) structure of the residual and a set of weights as in Newey and West (1987) have been used to take care of the problem of non-positive definite variance covariance matrix in finite samples.

[28] Other variables in the information set can be expected to be orthogonal to the error term of the Euler equation. However, it is a well-known result that the properties of the estimator can deteriorate if weak instruments are included. Hence, the relatively small set.

[29] An alternative to using mean expenditure as a benchmark is to take into account the cross-sectional sample heterogeneity of consumption growth. In this case, the estimated coefficient of risk aversion for the whole sample would be negative, when estimating the single-equation model for the return on the S&P500 CI or the two-equation model, and would be positive, but not significant, when using the return on T-bills.

consumption growth, as shown in equation (14).[30] A negative finding from the empirical analysis concerns the test of overidentifying restrictions, which surprisingly never rejects the theory restrictions, not even for predicted non-shareholders, nor for the whole sample of households.

6. Summary of Results and Concluding Remarks

This paper provides a unified framework of analysis to assess the role of portfolio heterogeneity in rationalizing the empirical inconsistency of the Consumption-based Capital Asset Pricing Model and shows that accounting explicitly for limited financial market participation helps to reconcile the theory with the empirical evidence. Specifically, I take the three best-known characterizations of the empirical failure of the model and assess the extent to which the empirical inconsistency depends on the use of aggregate data that ignore heterogeneity in financial market participation. The analysis shows that in all cases, relaxing the representative agent hypothesis, or at least pushing it to a lower level of aggregation, narrows, and in some cases eliminates completely the gap between theory and empirical evidence.

The first characterization of the empirical failure of the C-CAPM considered is Hansen and Jagannathan (1991). I assess the implications of limited financial market participation for the consistency of the IMRS with the volatility bounds that restrict the region for admissible stochastic discount factors. This exercise is extremely general and weaker than those based on the calibration and estimation of the model. In fact, it does not propose and test alternative parametric specifications but simply characterizes the set of stochastic pricing factors consistent with observed returns and verifies the consistency of the IMRS with this set. Under the hypotheses of isoelastic utility and of representative agent, which ensures that the IMRS is a valid discount factor, the IMRS is consistent with the bounds for relatively large values of relative risk aversion. Using disaggregated data, I succeed in moving the representative agent assumption to a lower level of aggregation and average over financial

[30] If one were able to model the conditional second moments of asset returns, one could impose additional restrictions in the estimation that would help assess further the ability of the model to explain the observed premiums.

market participants and non-participants separately. A plot of the IMRSs based on the representative shareholder and on the representative non-shareholder shows that for low risk aversion the former is closer to the bound than the latter, which is preliminary evidence that accounting for limited financial market participation can be important in assessing the empirical performance of the C-CAPM. Evaluating the role of portfolio heterogeneity in the presence of habit persistence or Epstein and Zin preferences is not straightforward.

The second set of tests is based on the calibration of the C-CAPM and verifies the implications of portfolio heterogeneity for the unconditional restrictions that the model implies. This exercise focuses explicitly on the relationship between the time-series properties of consumption and the coefficient of relative risk aversion. The results obtained are fully consistent with those obtained within Hansen and Jagannathan's framework but have the additional advantage of allowing somewhat more precise quantitative statements. They show unequivocally that shareholders' expenditure is more volatile and more closely correlated with asset returns than non-shareholders', and this helps explain why the risk aversion coefficient that reconciles their consumption profile with observed return is substantially lower than that of the representative agent. Further, although the point estimates obtained are still rather high, especially those based on the log-linear version of the model, shareholders' risk aversion turns out to be substantially lower than non-shareholders'.

The last type of test, more structural in approach, consists in the GMM estimation of the parameters of interest, which allows the use of information on the changes in expectation on consumption and asset returns, thus improving the efficiency of the estimates. The point estimates imply that, under the assumption of isoelastic utility, predicted shareholders' risk aversion is 7 to 8. The estimates for non-shareholders and for the sample representative agent are often negative or much higher, and they lack precision. In an important negative finding, the test of the overidentifying restrictions does not reject the null hypothesis of correct specification even for non-shareholders.

Overall, the evidence presented here suggests unequivocally that there are important differences between shareholders and non-shareholders. Ignoring these differences is likely to result in an upward biased estimate of relative risk aversion. Interestingly, on the one hand, the evidence for non-shareholders and the representative agent is not robust, and the results differ somewhat across exercises and in many cases lack economic meaning. On the other hand, the tests I have carried out point towards a value of risk aversion for shareholders around 7 or 8. The question is what a reasonable value for risk aversion is. Mehra and Prescott (1985) quote several micro-econometric estimates bounding risk aversion from above by 3, although there is some evidence in Kocherlakota (1990) that these estimates may be severely downward-biased. In their 1991 study, Mankiw and Zeldes show that an agent with risk aversion above 10 would be willing to pay unrealistically high amounts in order to avoid all risk. A minority of other economists believe that agents are more risk-averse than is commonly thought, and that levels of risk aversion higher than 10 imply absolutely reasonable behavior as long as the risk affects a limited fraction of total wealth. For example, Kandel and Stambaugh (1991) show that the amount that agents with a relative risk aversion of 30 are willing to pay to avoid a risk involving the loss of at most 1 % of their wealth is absolutely reasonable. However, this position is not common in the literature, and most economists believe that a risk aversion above 10, or even 5, implies totally unrealistic conduct.

Tables and Figures

Table 1

ASSET OWNERSHIP OVER TIME

Year	Percentage of households with: No financial assets	Checking accounts	Savings accounts	Checking and savings accounts	Checking/savings accounts and shares
1982	0.161	0.105	0.076	0.364	0.292
1983	0.155	0.128	0.074	0.367	0.273
1984	0.163	0.125	0.078	0.348	0.283
1985	0.148	0.135	0.065	0.350	0.298
1986	0.162	0.138	0.058	0.327	0.312
1987	0.151	0.123	0.064	0.329	0.328
1988	0.158	0.140	0.064	0.316	0.320
1989	0.158	0.133	0.053	0.328	0.321
1990	0.137	0.139	0.065	0.343	0.312
1991	0.152	0.135	0.052	0.328	0.328
1992	0.153	0.145	0.057	0.308	0.332
1993	0.156	0.144	0.057	0.296	0.344
1994	0.139	0.158	0.047	0.321	0.328
1995	0.151	0.141	0.053	0.303	0.345
1996	0.155	0.118	0.054	0.321	0.348
Total	0.153	0.135	0.061	0.330	0.316

NOTE: 1996 refers only to the first quarter.

Table 2

HOUSEHOLD CHARACTERISTICS

	Households with:			Total sample
	No financial assets	Checking/savings accounts	Checking/savings accounts and shares	
Age of head	42.845	43.494	45.706	44.080
Less than high school*	0.497	0.186	0.072	0.198
High school diploma*	0.454	0.587	0.518	0.545
College degree*	0.049	0.227	0.410	0.258
Male head*	0.491	0.665	0.755	0.667
White*	0.631	0.860	0.929	0.846
Black*	0.332	0.097	0.038	0.115
Asian*	0.019	0.018	0.012	0.016
With children*	0.555	0.401	0.428	0.434
Number of children	1.243	0.761	0.775	0.840
	(1.458)	(1.128)	(1.060)	(1.178)
Household size	3.104	2.664	2.796	2.774
	(1.912)	(1.561)	(1.405)	(1.582)
Single person*	0.235	0.272	0.186	0.239
North-East*	0.199	0.178	0.209	0.191
Mid-West*	0.216	0.241	0.259	0.243
South*	0.375	0.290	0.272	0.297
West*	0.210	0.291	0.260	0.268
After tax income (median)	$8,713	$19,314	$30,653	$20,586
Monthly consumption (median)	$426	$641	$856	$668
Quarterly consumption growth (median)	0.038	0.049	0.058	0.050
Quarterly consumption	0.083	0.100	0.112	0.101
growth (mean)	(0.316)	(0.318)	(0.314)	(0.317)
Quarterly per-capita	-0.005	0.002	0.002	0.001
consumpt. growth (mean)	(0.061)	(0.045)	(0.054)	(0.041)
No. of households	5,206	17,933	10,747	34,033

NOTE: * denotes "percentage of households". Standard deviations in parentheses. After-tax income is the median of household annual income as reported at the last interview. Monthly consumption is the median of household average monthly consumption. Quarterly consumption growth is the median and the mean of average household quarterly consumption growth over the year they participate in the survey. Quarterly per capita consumption growth denotes the median of overlapping rates of growth of the representative agent (mean) consumption. Income and expenditure are in 1982-1984 US dollars.

Table 3

RESULTS OF PROBIT ESTIMATION
(probability of holding risky assets)

	Parameter	Marginal Effect
Age	4.952	1.691
	(1.909)	(0.652)
Age^2	-7.799	-2.663
	(4.182)	(1.428)
Age^3	6.187	2.113
	(2.892)	(0.988)
High school diploma	0.786	0.259
	(0.089)	(0.028)
College degree	1.177	0.430
	(0.098)	(0.034)
Age*High school diploma	-0.239	-0.082
	(0.163)	(0.056)
Age*College degree	-0.314	-0.107
	(0.188)	(0.064)
Children (dummy)	0.024	0.008
	(0.020)	(0.007)
Single-person household	-0.094	-0.032
	(0.022)	(0.007)
Male	0.120	0.040
	(0.018)	(0.006)
North-East	0.040	0.014
	(0.023)	(0.008)
Mid-West	-0.083	-0.028
	(0.022)	(0.007)
South	-0.145	-0.049
	(0.023)	(0.007)
Black	-0.590	-0.172
	(0.030)	(0.007)
Asian	-0.269	-0.085
	(0.064)	(0.018)
Other non-white	-0.308	-0.096
	(0.053)	(0.015)
Earnings ($thousand)	0.015	0.005
	(0.000)	(0.000)
Constant	-2.700	-
	(0.281)	
p-value year dummies	0.0000	
No. of observations	34,033	
Pseudo R^2	0.1465	

NOTE: Standard errors in parentheses.

Table 4

DESCRIPTIVE STATISTICS FOR ASSET RETURNS
(1982:1 – 1996:1)

	Mean	Standard deviation	Min.	Max.
S&P500 CI	0.041	0.074	-0.295	0.267
T-bills	0.016	0.006	0.007	0.032
Equity premium	0.025	0.074	-0.309	0.249

NOTE: Quarterly returns on the Standard & Poor 500 Composite Index and on 3-month Treasury bills. The equity premium corresponds to the difference between the two returns.

Tab. 5

CALIBRATION OF THE RISK AVERSION PARAMETER
(the non-linear model)

	Isoelastic utility		Habit (δ=0.15)		Epstein and Zin utility ρ=0.5 (e.s.=2)		ρ=5 (e.s.=0.2)	
	e	γ	e	γ	e	γ	e	γ
Representative shareholders	0.0230 (0.0121)	7	0.2933 (0.1889)	7	2.9147 (1.5294)	8.5	300.343 (152.875)	59
Representative non-shareholders	0.0217 (0.0111)	13.5	-0.3725 (0.4556)	12	2.8189 (1.5739)	8	133.2103 (68.636)	67
Representative agent	0.0217 (0.0111)	9	1.1593 (0.8997)	10.5	2.5918 (1.3495)	8	56.5779 (29.0602	62

NOTE: Standard errors in parentheses. δ denotes the degree of persistence of habit. ρ is the inverse of the elasticity of substitution (e.s.). The quarterly discount rate is set equal to 2 %.

Tab. 6

CALIBRATION OF THE RELATIVE RISK AVERSION COEFFICIENT
(the log-linear model)

	(1)	(2)	γ
Shareholders	0.1428	0.0535	44
Non-shareholders	0.1213	0.0304	92
Representative agent	0.1238	0.0452	60

NOTE: Column (1) and (2) display the correlation between consumption growth and the return on the S&P500CI and the variance of consumption growth, respectively. For predicted shareholders and non-shareholders, they are given by $\hat{Corr}\left(\overline{\Delta \log c_{t+3}^{g}}, \log R_{t+3}^{S\&P500}\right)$ and $\hat{Var}\left(\overline{\Delta \log c_{t+3}^{g}}\right)$. For the representative agent, they are computed as $\hat{Corr}\left(\Delta \log \bar{c}_{t+3}, \log R_{t+3}^{S\&P500}\right)$ and $\hat{Var}\left(\Delta \log \bar{c}_{t+3}\right)$. The numerator of equation (17) in the text is: $\log \overline{R^{S\&P500}} - \log \overline{R^{Tbill}} = 0.0250$.

Tab. 7

GMM ESTIMATION OF THE EULER EQUATION

		Predicted shareholders	Predicted non-sharehold.	Representative agent
S&P500	γ^{-1}	0.1204	-0.0718	0.0806
		(0.0579)	(0.0400)	(0.0715)
	Constant	-0.0007	0.0054	-0.0031
		(0.0033)	(0.0027)	(0.0039)
	Sargan test	13.6596	10.3156	13.1998
	(p-value)	(0.5515)	(0.7994)	(0.5869)
T-bills	γ^{-1}	0.7402	0.1295	0.1923
		(0.3251)	(0.3639)	(0.3649)
	Constant	-0.0068	0.0005	-0.0031
		(0.0063)	(0.0059)	(0.0067)
	Sargan test	16.2366	10.3451	9.6882
	(p-value)	(0.3665)	(0.7975)	(0.8389)
S&P500 and T-bills	γ^{-1}	0.1326	-0.0700	0.0913
		(0.0571)	(0.0397)	(0.0708)
	Constant[a]	-0.0013	0.0053	-0.0035
		(0.0033)	(0.0027)	(0.0039)
	Constant[b]	0.0031	0.0034	-0.0015
		(0.0015)	(0.0022)	(0.0023)
	Sargan test	29.8056	20.8522	22.9205
	(p-value)	(0.5274)	(0.9158)	(0.8520)

NOTE: Standard errors in parentheses. Constant[a] and Constant[b] refer to the constant terms in the equations with the return on the S&P500 CI and on T-bills, respectively, when they are estimated jointly. The set of instruments includes: consumption growth lagged three and four quarters; the returns on T-bills and on the S&P500 CI and the rate of inflation lagged one to four quarters; and the risk and term spreads lagged one quarter. The degrees of freedom to the Sargan test of overidentifying restrictions are 15 in the single equation model and 31 in the two-equation model.

Fig. 1

QUARTERLY RETURN TO S&P500 CI AND TO 3-MONTH TREASURY BILLS
(1981:1-1996:1)

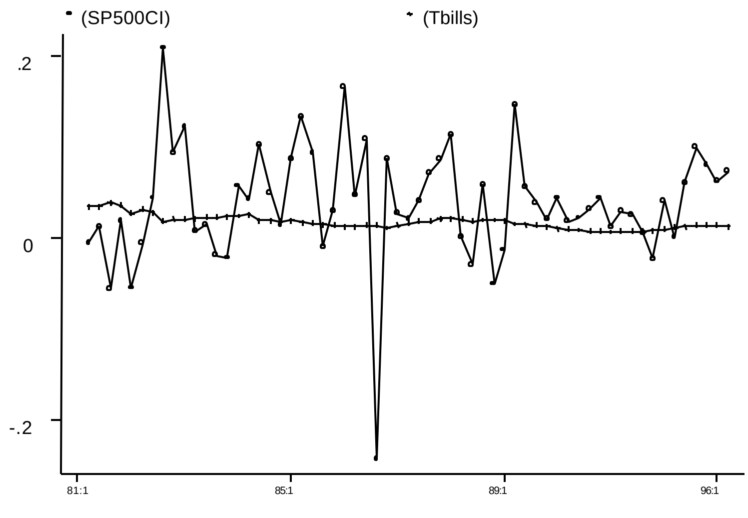
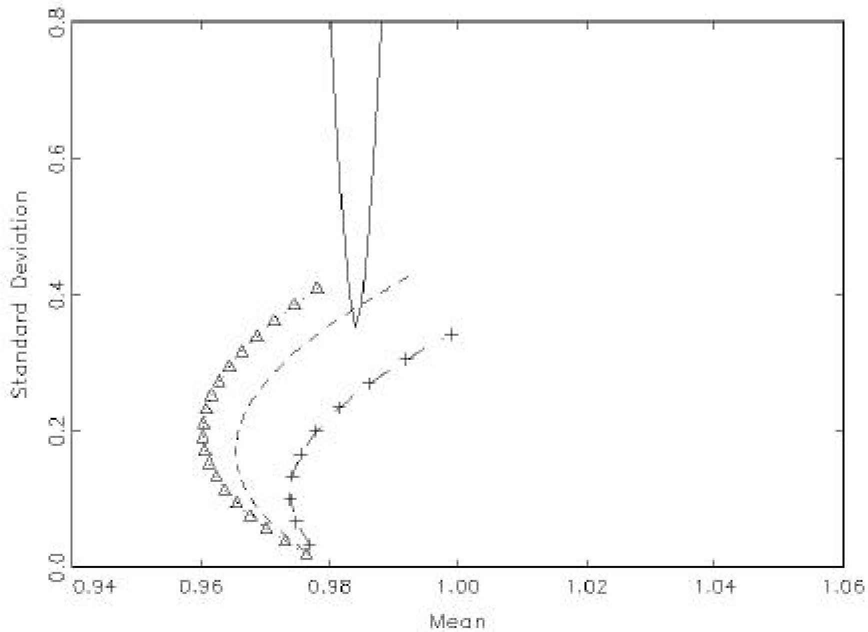


Fig. 2

HANSEN AND JAGANNATHAN BOUNDS:
THE CASE OF ISOELASTIC UTILITY

(Quarterly CEX expenditure on non-durables and services. S&P500 CI and T-bill returns, 1982:1-1996:1. Quarterly discount rate of 2 %. Relative risk aversion set to 0.5, 1, 1.5, etc. + and Δ denote the sequence for shareholders and non-shareholders, respectively. The dashed line refers to the representative agent of the sample.)

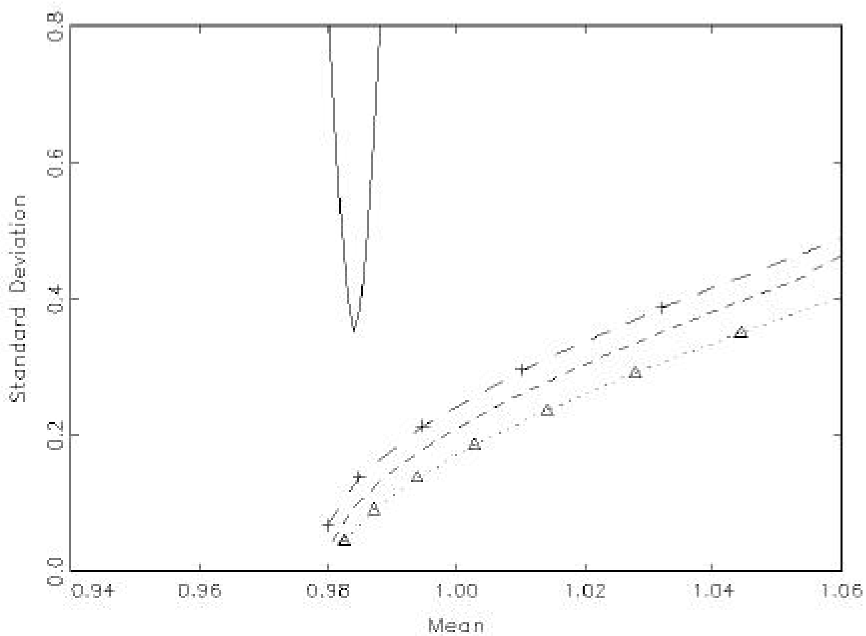


+ = predicted shareholders
Δ= predicted non-shareholders
- = representative agent

Fig. 3

HANSEN AND JAGANNATHAN BOUNDS:
THE CASE OF HABIT PERSISTENCE

(Quarterly CEX expenditure on non-durables and services. S&P500 CI and T-bill returns, 1982:1-1996:1. Quarterly discount rate of 2 %. Persistence parameter of 0.15. γ=0.5, 1, 1.5, etc. + and Δ denote the sequence for shareholders and non-shareholders, respectively. The dashed line refers to the representative agent of the sample.)

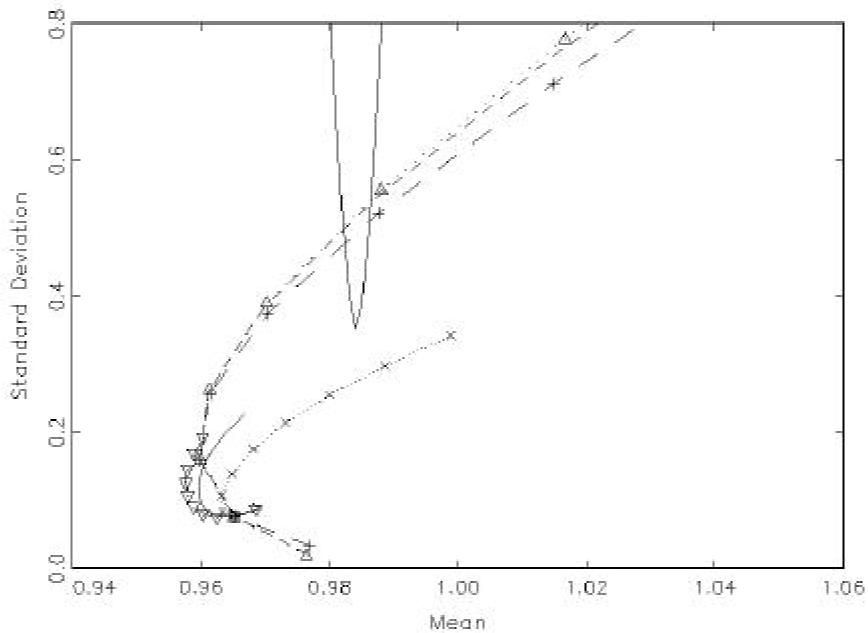


+ = predicted shareholders
Δ= predicted non-shareholders
- = representative agent

Fig. 4

HANSEN AND JAGANNATHAN BOUNDS:
THE CASE OF EPSTEIN AND ZIN UTILITY

(Quarterly CEX expenditure on non-durables and services. S&P500 CI and T-bill returns, 1982:1-1996:1. Quarterly discount rate of 2 %. Elasticity of intertemporal substitution of 2 (Δ and +) and 0.2 (∇ and ×). γ=0.5, 1, 1.5, etc. Δ and ∇ denote the sequences for shareholders; + and × denote the sequences for non-shareholders. The dashed lines refer to the representative agent of the sample.)



+ = predicted shareholders with elasticity of substitution equal to 2
Δ= predicted non-shareholders with elasticity of substitution equal to 2
- = representative agent with elasticity of substitution equal to 2

× = predicted shareholders with elasticity of substitution equal to 0.2
∇= predicted non-shareholders with elasticity of substitution equal to 0.2
— = representative agent with elasticity of substitution equal to 0.2

Appendix I

THE US CONSUMER EXPENDITURE DATA

The data used for the analysis are taken from the US Consumer Expenditure Survey (CEX), which is a representative sample of the US population, conducted by the US Bureau of Labor Statistics. The survey is a rotating panel, with each consumer unit being interviewed every three months over a twelve-month period, apart from attrition. About 4500 households are interviewed each quarter, more or less evenly spread over the three months: 80% are re-interviewed after three months, whereas the remaining 20% are dropped and a new group is added. So approximately one fifth of the units interviewed each month are new to the survey. This rotating procedure is designed to improve the overall efficiency of the survey and to attenuate attrition.

The data used for the analysis cover the period from 1982 to 1996, first quarter.[31] I exclude those households with incomplete income responses and those living in rural areas or in university housing. In addition, I exclude those whose head was under 21 or over 75 years old, those that do not participate in the last interview when the information on financial wealth is collected (about 30% of the initial sample), and those whose financial supplement contains invalid blanks in the stocks of assets held (around 20% of respondents). Finally, I drop those households with no consecutive interviews or reporting zero food expenditure and those whose total consumption on non-durables and services changes by more than 300% between contacts. As so restricted, the sample consists of 121,777 observations on 34,033 households. Given the timing of the observations on expenditure, which refer to the month preceding the one when the interview takes place, each household contributes at most four observations on monthly consumption at quarterly intervals. Overall, my sample spans over 171 months, with consumption observations available from 1981:12 to 1996:2.

[31] In 1986, the sample design and the household identification numbers were changed. For the first quarter of 1986, the Bureau of Labor Statistics created two files: one based on the original sample design and one based on the new design. After the first quarter, no track is kept of the households in the old sample. Since the distinction between shareholders and non-shareholders is based on information collected during the last interview only, those households whose first interview was in the third or fourth quarter of 1985 are excluded. For similar reasons, those whose first interview is after June 1995 are also excluded. Thus, the sample used consists of households whose first interview was between 1982:1 and 1985:6 or between 1986:2 and 1995:6.

The last column of Table 2 reports some descriptive statistics for the sample as whole. All figures are in 1982-1984 US dollars. Median annual household income after taxes is $20,600; median monthly consumption is $668. Median quarterly consumption growth is 0.05. The average age of the head is 44. 20% of heads do not have a high school diploma; 55% have a high school diploma; 26% have a college degree. 67% of household heads are a male; 85% are white; 12% are black and 24% of households are single persons. 43% of households have children. 19% live in the North-East, 24% in the Mid-West, 30% in the South and 27% in the West.

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